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O'Connor, R.C. and Dixon, D. and Rasmussen, Susan (2009) *The structure and temporal stability of the Child and Adolescent Perfectionism Scale*. Psychological Assessment, 21 (3). pp. 437-443. ISSN 1040-3590

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# The Structure and Temporal Stability of the Child and Adolescent Perfectionism Scale

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In this study, the authors examined the factor structure and temporal stability of the Child and Adolescent Perfectionism Scale (CAPS; G. L. Flett, P. L. Hewitt, D. J. Boucher, L. A. Davidson, & Y. Munro, 1997) in 2 samples of adolescents (15–16 years old). In Sample 1 ( $n = 624$ ), confirmatory factor analysis did not support a 2-factor structure (self-oriented and socially prescribed perfectionism). As in B. T. McCreary, T. E. Joiner, N. B. Schmidt, & N. S. Ialongo (2004), reanalysis suggested a 3-factor solution (i.e., socially prescribed perfectionism, self-oriented–Striving perfectionism, self-oriented–Critical perfectionism). The authors validated their 3-factor model in an independent replication sample (Sample 2;  $n = 514$ ) and confirmed that the 3-factor structure was invariant across gender and time (test–retest over 6 months). Taking these analyses together, the authors concluded that their discriminant 3-factor structure is robust. Theoretical and clinical implications are discussed. More research on the predictive validity of the CAPS is suggested.

*Keywords:* perfectionism, adolescent, child, factor structure, reliability

There is growing evidence that perfectionism is associated with psychological distress in clinical and nonclinical populations (see Flett & Hewitt, 2002; O'Connor & Sheehy, 2001; O'Connor, et al., 2007; Shafran & Mansell, 2001; see also Stoeber & Otto, 2006). However, compared with the adult literature, relatively few studies have involved investigation of the relationship between perfectionism and psychological health in children and adolescents (O'Connor, 2007; Rice & Preusser, 2002). This disparity may be due, in large part, to the relative lack of availability of specifically tailored child and adolescent perfectionism scales with published psychometric properties. For adults, there are a number of different measures of perfectionism that are reliable and valid (Enns & Cox, 2002). The most widely used of these are the two Multidimensional Perfectionism Scales (MPS) developed by Hewitt and Flett (1991) and by Frost, Marten, Lahart, and Rosenblate (1990), respectively. Although these research groups have identified different dimensions of perfectionism, they each posit that perfectionism is best conceptualized as having both personal and social components. We focus on Hewitt and Flett's (1991) measure, as it is the basis for the measure used herein.

Hewitt and Flett's (1991) scale comprises three dimensions: (a) *Self-oriented perfectionism* (SOP) is defined as a strong motivation to be perfect, with all-or-nothing thinking and self-reported high achievement expectations; (b) *socially prescribed perfectionism*

(SPP) assesses the extent to which an individual believes that others hold unrealistically high expectations of their behavior; and (c) *other-oriented perfectionism* (OOP) is the degree to which an individual sets unrealistic standards for others.

## Child and Adolescent Perfectionism

There are very few multidimensional perfectionism scales devised specifically for use with children and adolescents.<sup>1</sup> These scales include the Adaptive/Maladaptive Perfectionism Scale (Rice & Preusser, 2002) and the Child and Adolescent Perfectionism Scale (CAPS; Flett, Hewitt, Boucher, Davidson, & Munro, 1997).<sup>2</sup> Although the Adaptive/Maladaptive Perfectionism Scale has recently been validated (Rice, Kubal, & Preusser, 2004), to date, it has been used less frequently than has the Flett and colleagues' (1997) measure. The CAPS was developed from Hewitt and Flett's original MPS (Hewitt, Flett, Turnbull-Donovan, & Mikail, 1991; Hewitt, Newton, Flett, & Callander, 1997). In addition to being composed of different items, the CAPS is further distinguished from the MPS because it comprises two subscales rather than three: SPP (10 items, e.g., "There are people in my life who expect me to be perfect") and SOP (12 items; e.g., "I try to be perfect in everything I do"). From those studies that have used the CAPS, there is evidence that perfectionism is associated with psychological distress and maladjustment in children and adoles-

cence (Boergers, Spirito & Donaldson, 1998; Castro-Fornieles et al., 2007; Donaldson, Spirito, & Farnett, 2000; Hewitt, Caelian, Flett, Sherry, Collins, & Flynn, 2002; Hewitt et al., 1997; O'Connor, Rasmussen, & Hawton, in press).

Although the two-factor structure and the reliability of the CAPS were reported at a Canadian conference (Flett, Hewitt, & Davidson, 1990), with two exceptions (Castro et al., 2004; McCreary et al., 2004), studies investigating the psychometric properties of the CAPS have not been published elsewhere. Castro et al. (2004) reported that the CAPS had good internal consistency properties and adequate 1 week test-retest reliability. The only study to investigate the factor structure of the CAPS (McCreary et al., 2004), in a sample of 11–12-year-old African American school children, did not find sufficient evidence for the adequacy of Flett et al.'s (1997) two-factor structure. These authors concluded that a three-factor structure was a much better fit for the data (following item exclusion, the three factors were derived from 14 of the 22 items of the CAPS). In McCreary et al.'s (2004) study, SPP emerged as a single factor, although the original SOP items were better modeled as two factors; SOP-Striving (defined as striving to perfectionism) and SOP-Critical (defined as self-criticism). This finding is consistent with other theoretical considerations of the concept of perfectionism and with empirical evidence that SOP is not homogeneous (Dunkley, Blankstein, Masheb & Grilo, 2006; Hunter & O'Connor, 2003; O'Connor, 2007; O'Connor & O'Connor, 2003). Rather it may be better represented as having maladaptive and adaptive components.<sup>3</sup> Indeed, there is a growing consensus that striving for high standards loads on an adaptive, higher order factor entitled Personal Standards perfectionism, whereas critical type items load onto a maladaptive, higher order factor, Evaluative Concerns perfectionism (Dunkley et al., 2006; see also Hewitt, Flett, Besser, Sherry & McGee, 2003; O'Connor, 2007).

In the light of these conflicting factor solutions (i.e., two factors vs. three factors), the central aim in the present study was to investigate the factor structure of the CAPS in two samples of adolescents by examining Flett et al.'s (1997) and McCreary et al.'s (2004) two- and three-factor structures, respectively. In addition, as gender differences in perfectionism have been reported elsewhere (Donaldson et al., 2000; Hewitt, Flett, & Turnbull-Donovan, 1992; McCreary et al., 2004) and personality dimensions are usually stable in the short-term (e.g., Fullana et al., 2007; Hewitt & Flett, 1991; Hewitt et al., 1992), we also investigated measurement invariance in boys and girls and across time (6 months). Furthermore, a recent systematic review suggested that testing for gender differences should be conducted as a matter of course when examining perfectionism and when prospective studies are lacking (O'Connor, 2007). In addition, for maximization of clinical and educational usefulness, it is important to determine whether the CAPS assesses perfectionism reliably over time.

## Participants, Measures, and Procedure

### Sample 1

We recruited 624 adolescents from schools in Scotland. Very few adolescents who were invited to participate declined. Approximately 80% of those school pupils eligible to take part (in Samples 1 and 2) completed the measures. Nonparticipation was due largely to timetable and logistical issues that precluded partaking.

There were 322 girls and 299 boys (3 respondents did not indicate gender) with an overall mean age of 15.6 years ( $SD = 0.9$ ). Of the participants, 95% were White, 4% were Asian, and 1% were of another ethnic group. We felt that 15–16 years was an appropriate age group to test the structure of the CAPS, given self-consciousness increases during adolescence and given that the “impact of socially prescribed pressures to be perfect are magnified substantially during adolescence” (Flett, Hewitt, Oliver, & Macdonald, 2002, p. 115).

### Sample 2

We recruited a new sample of 737 adolescents. There were 367 girls and 369 boys with an overall mean age of 15.2 years ( $SD = 0.7$ ). Of the participants, 95% were White, 4% were Asian, and 1% were of another ethnic group. At Time 1 and Time 2, 6 months later, participants completed the CAPS. At Time 1, we recruited 737 respondents, and at Time 2, we followed up with 514 of these young people, thereby yielding a response rate of 70%. Although there were no gender differences, those who completed measures at both time points were significantly younger than were noncompleters ( $M = 15.13$  years,  $SD = 0.69$  vs.  $M = 15.36$  years,  $SD = 0.74$ )  $t(735) = 3.95$ ,  $p < .001$ , Cohen's  $d = .32$ . Those who completed the follow-up also did not differ from those who did not complete the follow-up in SPP,  $t(735) = 0.38$ , *ns*, and in SOP-Critical,  $t(735) = 1.36$ , *ns*, but did report higher levels of SOP-Striving ( $M = 10.8$ ,  $SD = 2.6$ ),  $t(735) = 4.94$ ,  $p < .001$  than did the noncompleters ( $M = 9.7$ ,  $SD = 2.9$ ).

We obtained ethical approval from the university's psychology department ethics committee. In Sample 2, to ensure anonymity but to allow for follow-up, respondents were asked to answer a series of questions (e.g., “Please write in the last two letters of your home postcode”) at both time points, which generated a unique reference code. There were no duplicate codes. All participants completed the 22-item Child and Adolescent Perfectionism Scale (CAPS; Flett et al., 1997). Respondents rated each statement on a 5-point Likert-type scale, ranging from 1 (*false—not at all true of me*) to 5 (*very true of me*). The SPP and SOP have been reported to be internally consistent (.86 and .85, respectively) and to be reliable over a 1 week period (Castro et al., 2004), and they are associated with depression and anxiety (Hewitt et al., 2002). SPP is also associated with suicide ideation (Hewitt et al., 1997) and self-harm (O'Connor, Rasmussen, Miles, & Hawton, 2009). Items 3, 9, and 18 from the CAPS were reverse scored to ensure that a higher score indicated greater perfectionism for all items.

### Factor Structures Tested

We tested Flett et al.'s (1997) and McCreary and colleagues' (2004) two- and three-factor structures, respectively. Therefore, the two-factor structure refers to SPP and SOP, and the three-factor structure refers to SPP, SOP-Critical, and SOP-Striving.

### Statistical Analyses

Data were analyzed with confirmatory factor analyses (CFA), with the EQS (6.1) software package (Bentler, 2004). Models were

<sup>3</sup> It is worth noting that in the perfectionism literature, SPP usually loads on the maladaptive factor, and SOP loads on the adaptive factor.

estimated with covariance matrices and maximum likelihood (ML) procedures. All variables showed acceptable levels of univariate skew and kurtosis. However, multivariate kurtosis was evident; consequently, the robust correction procedure for nonnormal data was applied throughout (Satorra & Bentler, 1994).

### *Assessment of Fit*

Three<sup>4</sup> fit indices were used, namely, the root-mean-square error of approximation (RMSEA; Browne & Cudeck, 1993), the comparative fit index (CFI) and the nonnormed fit index (NNFI). A RMSEA value of .07 or below indicates an acceptable fit, and a good fit is indicated by a value of .05 and by whether the whole of the 90% confidence interval (CI) for the statistic falls below .06; traditionally, NNFI and CFI values of >.90 indicate adequate fit, however, values closer to .95 are preferred, and values  $\geq$ .95 indicate good fit (Hu & Bentler, 1999). In the current study, models with both NNFI and CFI values of .95 or above are described as having good fit; values between .92 and .94 are described as having adequate model fit, and models with either NNFI or CFI values between .91 and .90 are described as having marginal fit.  $R^2$  values indicated the strength of the relationship between the target latent construct and each measurement item. The Lagrange multiplier (LM) test was applied to the factor loadings and error covariances, to investigate the effect of freeing specific parameters on the fit indices. Nonnested models were compared with Akaike's information criterion (AIC) values (Akaike, 1987). The model with the smallest AIC value is viewed as being the more parsimonious and better fitting model.

### *Validation Testing*

The factor structure of the CAPS, generated with Sample 1, was validated against an independent sample of schoolchildren, Sample 2, with a multigroup invariance testing protocol (Byrne, 2006). The model specification for Sample 2 (the validation sample) was identical to that of Sample 1 (the baseline sample), including all specified start values derived from Sample 1. All factor loadings and error covariances in Sample 2 were constrained to be equal to those of Sample 1. Noninvariance was evidenced by the LM test of equality constraints; constrained parameters with univariate, incremental chi-square values with  $p < .05$  were viewed as noninvariant. The fit indices for multigroup invariance testing refer to model fit across both groups simultaneously.

### *Invariance Testing*

Invariance in the measurement model was tested across gender in both samples and across time in Sample 2. A hierarchical approach to invariance testing was used. First, baseline models for each group (boys and girls; Time 1 and Time 2) were developed separately. Second, the multigroup representation of the baseline models was tested for goodness of fit. In this configural model, parameters were not constrained to be equal across groups. This configural model functioned as the baseline model against which subsequent invariant models were compared. The fit indices refer to overall model fit across both samples simultaneously, not the fit of each baseline model individually. Third, measurement invariance was tested. All factor loadings and error covariances common

to both baseline models were constrained to be equal across the two groups. The LM test of equality constraints, described above, identified parameters that were noninvariant.

## **Results**

### *Examination of the Factor Structure of the Child and Adolescent Perfectionism Scale*

In the first instance, we examined Flett et al.'s (1997) two-factor solution for the CAPS, in which 10 items indicate the SPP factor and 12 items indicate the SOP factor. None of the criteria for acceptable model fit were met (NNFI = .81; CFI = .83; RMSEA = .083 [CI = .078–.088]). Low  $R^2$  values identified items that were poor indicators of their target factor; the three negatively worded items had  $R^2$  values below .05, and two of the three had nonsignificant path coefficients. These observations are consistent with previous studies that have shown negatively worded questionnaire items to be poor indicators of their target factors (D. Dixon, Johnston, Rowley, & Pollard, in press). The three negatively worded items were removed from all subsequent analyses of the two-factor model. Following removal of the three negatively worded items, the two-factor model was reestimated; the model continued to be inadequate (NNFI = .84; CFI = .86; RMSEA = .087 [CI = .082–.093]). The LM test continued to indicate improvements in model fit would be achieved with the inclusion of (a) six error covariances and (b) cross-loadings for two indicators of the SPP factor on the SOP factor. These changes could not be justified on theoretical grounds; therefore, we concluded that the two-factor solution was a poor representation of the data.

Next, we tested McCreary et al.'s (2004) three-factor structure for the CAPS. As this is a confirmatory analysis of an existing model, we tested this 14 item, three-factor model on Sample 1 with no modifications. The fit indices for this model were marginal (NNFI = .91; CFI = .92; RMSEA = .072 [CI = .064–.080]). Consequently, we returned to the full 22-items and examined whether a new three-factor model better represented the Sample 1 data. Each SOP item was labeled as either Critical or Striving, as indicated by McCreary et al. (2004) and consistent with O'Connor (2007).

The three-factor model was a poor fit (NNFI = .83; CFI = .85; RMSEA = .078 [CI = .073–.083]). Consistent with the two-factor model and McCreary et al. (2004), the three negatively worded items had  $R^2$  values below .05, and two of the three items had nonsignificant path coefficients. These items were removed from subsequent analyses. LM tests indicated significant improvements in model fit would be achieved if Item 16 ("When I do something, it has to be perfect") also indicated SOP-Critical and if Item 7 ("It really bothers me when I don't do my best all the time") also indicated SOP-Striving. Although these modifications improved model fit (NNFI = .90; CFI = .92; RMSEA = .068 [CI = .062–.074]), the fit indices remained only marginal. A series of post hoc modifications to the model were made, based on the LM tests. These modifications were the estimation of one error covariance (between Items 5 ["There are people in my life who expect

<sup>4</sup> Chi-square values were also calculated for all models. In all cases, the chi-square was significant, that is, it indicated inadequate fit; however, chi-square values are severely affected by large sample sizes. Consequently, three alternative fit indices are provided.



me to be perfect”) and 8 [“My family expects me to be perfect”], both indicators of SPP) and three additional cross-loadings (Item 15 [“People around me expect me to be great at everything”] indicated SPP and SOP-Striving, Items 4 [“I feel that I have to do my best all the time”] and 19 [“I am always expected to do better than others”] both indicated SOP-Striving and SPP). As Items 5 and 8 are semantically similar (i.e., expectations to be perfect), the error covariance is not unexpected. In addition, the cross-loading items may be conceptually mixed; for example, Item 19 suggests striving (“doing better”) and social comparison (“than others”). The reestimated model resulted in adequate fit indices (NNFI = .93; CFI = .94; RMSEA = .057 [CI = .051–.063]).

To maximize the discriminant validity of the three factors within the model, we removed cross-loading items (i.e., Items 4, 7, 15, 16, 19) and reestimated the model. This latter discriminant three-factor model yielded improved fit indices (NNFI = .95; CFI = .96; RMSEA = .057 [CI = .048–.065]). Comparison of the AIC values for the final two three-factor models indicated that the discriminant model was a more parsimonious representation of the data (AIC = 144.3 and 72.1 for the final three-factor and discriminant three-factor models, respectively). The items that comprised each factor are summarized in Table 1 (together with Cronbach’s alphas). Our solution had 11 of the 14 items in common with McCreary et al.’s (2004) three-factor solution. Items 12, 20, and 22 were not included in McCreary et al.’s solution, and Items 4 (“I feel that I have to do my best all the time”), 15 (“People around me expect me to be great at everything”), and 19 (“I am always expected to do better than others”), which appeared in the latter solution, did not enter our three-factor solution. To ensure that the discriminant three-factor model was not simply the result of capitalization of chance, we validated the model on an independent sample.

### Validation of the Discriminant Three-Factor Measurement Model

The discriminant three-factor model and the parameter estimates derived from the application of that model to Sample 1 (detailed above) were applied to Sample 2 ( $n = 514$ ). The model was an adequate fit to Sample 2 (NNFI = .93, CFI = .94, RMSEA = .058 [CI = .047–.067]). The model was then applied simultaneously to both samples to test whether the discriminant three-factor model replicated across Sample 2. The start values generated by Sample 1 were applied to both samples, and the parameter estimates were constrained to be equal across the two groups. The fit indices were adequate (NNFI = .94; CFI = .95; RMSEA = .057 [CI = .050–.063]). These data validate the discriminant three-factor model. However, of the 12 constraints applied in the validation process, 2 had univariate incremental chi-square  $ps$  of less than .05; these constraints were the path between SPP and Item 21 ( $p = .019$ ) and the path between SOP-Critical and Item 14 ( $p = .03$ ), indicating the strength of relationship between these items and their target factors was not invariant across the two samples. Nonetheless, the constraints imposed during this validation test were extremely rigorous (Byrne, 2006), and the validation did not indicate any conceptual differences; consequently, we conclude that the discriminant three-factor structure is robust and valid.

### Testing for Invariance in Factor Structure Across Gender

Baseline models were established separately for boys and girls in each sample. The discriminant three-factor model showed adequate fit for both boys (NNFI = .93; CFI = .94; RMSEA = .066 [CI = .052–.079]) and girls (NNFI = .95; CFI = .96; RMSEA = .052 [CI = .039–.065]) in Sample 1. The fit indices for the

Table 1

*The 14 Items of the Child and Adolescent Perfectionism Scale (CAPS) Which Compose the Shortened Discriminant Three-Factor CAPS-14*

Item	SOP-Striving $R^2$	SPP $R^2$	SOP-Critical $R^2$
1. I try to be perfect in everything I do	70		
2. I want to be the best at everything I do	53		
5. There are people in my life who expect me to be perfect		47	
6. I always try for the top score on a test	23		
8. My family expects me to be perfect		53	
10. People expect more from me than I am able to give		37	
11. I get mad at myself when I make a mistake			37
12. Other people think I have failed if I do not do my very best all the time		52	
13. Other people always expect me to be perfect		67	
14. I get upset if there is even one mistake in my work			54
17. My teachers expect my work to be perfect		26	
20. Even when I pass, I feel that I have failed if I didn't get one of the highest marks in the class			48
21. I feel that people ask too much of me		49	
22. I can't stand to be less than perfect			59
Sample 1 Cronbach's $\alpha$	.72	.85	.74
Sample 2 Time 1 Cronbach's $\alpha$	.72	.84	.72
Sample 2 Time 2 Cronbach's $\alpha$	.78	.86	.75

*Note.* SOP-Striving = self-oriented perfectionism–Striving; SPP = socially prescribed perfectionism; SOP-Critical = self-oriented perfectionism–Critical; CAPS-14 = 14 item Child and Adolescent Perfectionism Scale. The items are from “The child–adolescent perfectionism scale: Development, validation, and association with adjustment,” by G. L. Flett, P. L. Hewitt, D. J. Boucher, L. A. Davidson, and Y. Munro, 1997. Copyright 1997 by G. L. Flett, P. L. Hewitt, D. J. Boucher, L. A. Davidson, and Y. Munro. Reprinted with permission.

configural model based on the two baseline models were adequate (NNFI = .94; CFI = .95; RMSEA = .059 [CI = .050–.069]). Invariance in the measurement model was then assessed. The model showed good measurement invariance with adequate fit indices (NNFI = .94; CFI = .95; RMSEA = .058 [CI = .049–.067]). There was evidence of noninvariance in one parameter between the groups, namely, the path coefficient between Item 6 and SOP-Striving. This constraint had a marginally significant, incremental, univariate chi-square value ( $p = .047$ ). Consequently, we conclude that with the exception of a single factor loading, the discriminant three-factor model shows adequate measurement invariance across gender for Sample 1.

The model showed similar measurement invariance across gender in Sample 2. In Sample 2, the baseline model for the female group required the estimation of one error covariance between Item 13 and Item 10 (NNFI = .92; CFI = .94; RMSEA = .069 [CI = .054–.083]). The model did not require any modifications to achieve adequate fit to the male group (NNFI = .92; CFI = .94; RMSEA = .053 [CI = .037–.068]). The configural model showed adequate fit (NNFI = .92; CFI = .94; RMSEA = .061 [CI = .051–.072]). With the exception of the error covariance unique to the female sample, the model was then constrained to be equal across the two groups. The fit indices for the measurement invariance model were adequate (NNFI = .92; CFI = .93; RMSEA = .060 [CI = .049–.070]). The path between Item 14 and SOP-Critical was noninvariant across the groups (incremental univariate chi-square was significant for the release of this constraint,  $p = .013$ ). These data suggest partial measurement invariance across gender in Sample 2.

### *Testing for Invariance in Factor Structure Across Time*

The CAPS was measured, in Sample 2, at two time points, and measurement invariance of the discriminant three-factor model was assessed across this period. Baseline models for the Time 1 and Time 2 measures required the inclusion of one error covariance (between Item 21 and Item 10). The discriminant three-factor model was a good fit at Time 1 (NNFI = .95; CFI = .96; RMSEA = .046 [CI = .035–.056]) and an adequate fit at Time 2 (NNFI = .93; CFI = .94; RMSEA = .062 [CI = .052–.071]). The configural model fitted the data adequately (NNFI = .94; CFI = .95; RMSEA = .054 [CI = .035–.061]). The model was constrained to be equal across the time points, including the additional error covariance, and the model displayed good fit (NNFI = .95; CFI = .95; RMSEA = .052 [CI = .045–.058]). All 13 constraints imposed on the model were invariant across the two time points.

### *Temporal Stability of CAPS*

To investigate the stability of respondents' responses in Sample 2 between Time 1 and Time 2, we calculated intraclass correlation coefficients for each of the discriminant CAPS factors between Time 1 and Time 2. They were all highly significant ( $p < .001$ ; SPP = .61; SOP-Critical = .65; SOP-Striving = .64). Although there were no significant Time  $\times$  Gender interactions for any of the factors, two-way analyses of variance (Time  $\times$  Gender) revealed that respondents reported significantly lower SPP scores at Time 2 ( $M = 17.97$ ;  $SD = 6.27$ ) than at Time 1 ( $M = 18.47$ ;  $SD = 6.09$ ),  $F(1, 512) = 4.42$ ,  $p < .05$ . However, the SPP Time 1–Time

2 difference was small (Cohen's  $d = .09$ ; Cohen, 1988; Morris & DeShon, 2002). In addition, respondents also reported significantly lower SOP-Striving scores at Time 2 ( $M = 10.17$ ;  $SD = 2.88$ ) than at Time 1 ( $M = 10.75$ ;  $SD = 2.63$ ),  $F(1, 512) = 31.95$ ,  $p < .001$ , and across both time points, boys ( $M = 10.93$ ;  $SD = 2.29$ ) reported significantly higher levels of SOP-Striving than did girls ( $M = 9.96$ ;  $SD = 2.63$ ),  $F(1, 512) = 20.21$ ,  $p < .001$ . The size of the differences between Time 1 and Time 2 (Cohen's  $d = .25$ ) and between boys and girls (Cohen's  $d = .39$ ) were small to medium effects. There were no significant differences across time,  $F(1, 512) = 2.12$ , *ns*, or gender for SOP-Critical,  $F(1, 512) = 2.68$ , *ns*.

## *Discussion*

The central aim of the present study was to investigate the factor structure of the CAPS in two independent samples of adolescents. Consistent with McCreary et al. (2004), our findings supported a three-factor structure for the CAPS (namely SPP, SOP-Striving and SOP-Critical), rather than the two-factor structure (i.e., SPP and SOP) posited by Flett and colleagues (Flett et al., 1997). However, our new three-factor structure (CAPS-14) was a better fit for the data than was McCreary et al.'s. We also validated our three-factor discriminant model rigorously in an independent replication sample, and we confirmed that the three-factor structure was invariant across gender and time (over 6 months). The minor differences in modifications required during invariance testing between samples and across time might be accounted for by the fact that the two samples comprised different subgroups of participants. Those participants who completed the CAPS at Time 1 and Time 2 (Sample 2) were significantly younger than were those who only completed the CAPS at Time 1 and significantly younger than were the Sample 1 participants. The confirmation that the SOP items resolve into two factors is consistent with recent research on its adaptive and maladaptive components (e.g., Adkins & Parker, 1996; Dunkley et al., 2006; Enns & Cox, 1999; O'Connor, 2007; O'Connor & O'Connor, 2003; Stoeber & Otto, 2006).

A major advantage of our 14-item measure (CAPS-14) is that it is a more parsimonious measure of perfectionism than is the original 22 item measure. Indeed, a recent review article on perfectionism concluded that further development of briefer versions of measurement scales would be beneficial (O'Connor, 2007). It is also noteworthy that the items in our measure are similar to those reported by McCreary et al. (2004) with 11 of the 14 items being common to both studies. There are a number of possible explanations for the minor differences between the two measures. First, the latter study comprised 11-year-olds and 12-year-olds, compared with 15-year-olds and 16-year-olds in the present study. As noted earlier, as self-consciousness develops throughout adolescence (Flett et al., 2002), the differences may be because the impact of perceived pressure varies as a function of age (Graham et al., 1987; Kandel & Andrews, 1987). Second, the vast majority of respondents in our samples were White Europeans, compared with African Americans in McCreary et al.'s (2004) study. Consequently, cultural influences require further exploration (McCreary et al., 2004; Nilsson, Paul, Lupini, & Tatem, 1999; Oyserman, Gant, & Ager, 1995). In addition, in the future, researchers should further investigate the relationship between adaptive and maladaptive perfectionism across cultures, given the equivocal differential

findings for adaptive perfectionism by ethnicity in the literature (Castro & Rice, 2003; Chang, Watkins, & Hudson Banks, 2004). Third, given that many of the individual items are significantly intercorrelated, there is probably some degree of item redundancy.

Although McCreary et al. (2004) and the present study yielded 11 items common to both samples, in future research, it could usefully be determined whether these items are core to the measurement of child and adolescent perfectionism and whether they are invariant across culture and age. In addition, given that the factor structure of the CAPS changes as a function of age and ethnic group, we would recommend that the original 22-item version be administered initially for groups that differ in composition from the present sample. Consistent with other temporal stability studies of personality (Fullana et al., 2007; Hewitt & Flett, 1991; Hewitt et al., 1992), we also yielded reasonable evidence of temporal stability over a 6 month period. Although respondents' scores on the SPP and SOP-Striving subscales were significantly lower at Time 2 than at Time 1, the effect sizes are small (Cohen, 1988). These differences are not accounted for by differences in sample composition between Time 1 and Time 2 because those who did, versus did not, complete measures at both time points did not differ in SPP. In addition, those who completed measures at both time points recorded higher SOP-Striving scores than did those who completed only the T1 measure. With the exception of SOP-Striving, SPP and SOP-Critical responses were similar for boys and girls. However, it is worth noting that the boys reported setting higher self-standards than did girls. It would be worth exploring this further to determine whether this gender difference is domain dependent, especially given the evidence that boys have higher performance expectations in mathematics than do girls, whereas the opposite is the case for language and verbal performance (e.g., Skaalvik & Skaalvik, 2004).

To our knowledge, this is the first study to both investigate the factor structure of the CAPS and to validate it in an independent sample over time. The analyses of the data from our two samples were rigorous and robust. Our findings suggest that a brief, 14-item measure of the CAPS is robust and largely temporally stable over 6 months. Nonetheless, in the future, researchers should endeavor to replicate our proposed factor structure in different age and ethnic groups.

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Received March 12, 2008

Revision received January 20, 2009

Accepted March 30, 2009 ■